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Horizontal inequities in Australia's mixed public/private health care system

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Abstract

Recent comparative evidence from OECD countries suggests that Australia's mixed public–private health system does a good job in ensuring high and fairly equal access to doctor, hospital and dental care services. This paper provides some further analysis of the same data from the Australian National Health Survey for 2001 to examine whether the general finding of horizontal equity remains when the full potential of the data is realized. We extend the common core cross-country comparative analysis by expanding the set of indicators used in the procedure of standardizing for health care need differences, by providing a separate analysis for the use for general practitioner and specialist care and by differentiating between admissions as public and private patients.

Overall, our analysis confirms that in 2001 Medicare largely did seem to be attaining an equitable distribution of health care access: Australians in need of care did get to see a doctor and to be admitted to a hospital. However, they were not equally likely to see the same doctor and to end up in the same hospital bed. As in other OECD countries, higher income Australians are more likely to consult a specialist, all else equal, while lower income patients are more likely to consult a general practitioner. The unequal distribution of private health insurance coverage by income contributes to the phenomenon that the better-off and the less well-off do not receive the same mix of services. There is a risk that – as in some other OECD countries – the principle of equal access for equal need may be further compromised by the future expansion of the private sector in secondary care services. To the extent that such inequalities in use may translate in inequalities in health outcomes, there may be some reason for concern.

Keywords: Equity; Australia; Access; Utilisation; Insurance system

1. Introduction

Australia provides universal coverage through a publicly financed health scheme, known generally as Medicare. It also has a significant private sector, with private expenditure constituting around 32% of total health expenditure, which is one of the highest

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proportions within the OECD [1]. Out-of-pocket spending on health care in Australia is higher than in many similar OECD countries and has been growing rapidly [1]. Free public hospital care and free or subsidized medical care should ensure that equitable access to health care, at least in terms of financial barriers, is achieved. However, there must be sufficient capacity in the public hospital system for those without private insurance to have the same access to treatment as those with insurance that provides for private hospital care. Private hospital use and private admissions to public hospitals have increased more rapidly than public admissions in public hospitals. Since 1996, government policy has encouraged the growth of private health insurance with subsidies and tax incentives. By 2001, the proportion of the population with private insurance had risen to 45% compared to 30% in 1998 [2]. So it is not clear that the entitlement to free hospital care is sufficient to deliver equitable access. Out of hospital medical care is subsidized according to a fee schedule set by Medicare but medical providers are free to set their own fee level and the patient meets any above schedule cost as an out-of-pocket payment when the fee is above the Medicare fee. The proportion of medical consultations for which no co-payment was charged was 70% in 2001, but this varies across type of medical practice with a higher bulk-billing in general practice. Unless there is sufficient price competition among providers, at least in low-income regions, co-payments will rise. Again the entitlement to free medical care is not sufficient to ensure equitable access. Since 2001 there has been increasing concern over the growth in co-payments and new strategies have been introduced to limit exposure to out-of-pocket costs. There have also been new initiatives further supporting the private health insurance industry. Consequently, it is important to know whether the universal public system has delivered equitable access to care, and to establish a base-line against which the equity impact of further policy changes can be assessed.

A recent five-country comparative study based on the 2001 Commonwealth Fund International Health Policy Survey compared inequities in access to medical care in Australia to those observed in Canada, New Zealand, United Kingdom, and United States [3]. On the basis of a multivariate analysis of a 2001 random telephone survey of 1400 adults in each of these countries, they conclude that here were few significant access differences by income in Australia; yet, compared to UK, Australians were more likely to report out-of-pocket costs, and those with above-average income were more likely to have private supplemental insurance which protected them from cost-related access problems. However, this study uses respondents' own ratings about whether they are in high or low income and whether or not they experienced any access problems.

In the health economics literature, a fairly standard approach to the testing and measurement of income-related inequity in actual health care utilisation using health interview surveys has been developed [21] and applied in a number of cross-country comparisons of income-related inequity [4–6]. An early study using similar methods was done for Australia [7] using the 1989–1990 National Health Survey. They find mixed results but this is largely due to their separate use of various morbidity variables as health care need indicators. Recently, Australia was included for the first time in a similar comparison involving 21 OECD countries and using a more appropriate and uniform methodology [8,9]. The Australian results show a slight pro-poor inequality in both the probability of hospital visit and no inequality in the probability of doctor visits over the last 12 months. While this suggests that the Australian system performs well on the equitable access criterion, this conclusion has to be treated cautiously. Data availability (from the Australian 2001 National Health Survey) allowed only a limited number of comparisons with other OECD countries. For example, the 2001 NHS asked for recall of whether there had been a GP or specialist doctor visit in the last 2 weeks, how many such visits and the time since the last consultation with either type of doctor. For hospital admissions, data were collected on number of admissions in the last 12 months and the number of nights and patient type at the most recent admission. A finding for other OECD countries was that while the majority had either pro-poor or non-significant levels of inequality in the overall number of doctor visits in the last 12 months, the distribution of general practitioner visit was generally pro-poor while the distribution of specialist visits was generally pro-rich. The 2001 NHS did not allow this comparison over the 12-month recall frame.

The aim of this paper is to explore further the issue of income-related use of health care in Australia. In particular, we examine inequality in hospital use split

by patient status, public or private, and the distribution of GP and specialist use over the 2-week time frame. We also undertake a decomposition of the inequality measure for GP and specialist visits to shed light on how different factors contribute to the inequalities in access to care. The horizontal version of the equity principle is interpreted to require that people in equal need of care be treated equally, irrespective of characteristics such as income, place of residence, race, etc. It is this principle of horizontal equity that the present study uses as the yardstick for its assessment of the Australian situation.

2. The Australian health care system and equity

Medicare covers basic hospital, medical services and pharmaceuticals. All Australians are entitled to be treated free of charge as public patients in public hospitals. Out of hospital medical services are for the most part provided by private medical practitioners on a fee for service basis. The patient is reimbursed by Medicare according to a fee schedule set by government (at 85% of schedule fees during this period). However, if the provider bills government directly (known in Australia as bulk billing), the Medicare reimbursement must be accepted as full payment, and the patient faces no out-of-pocket cost. The government also subsidizes drugs listed on the Pharmaceutical Benefits Scheme. Co-payments for prescriptions do not vary by drug but concession cardholders, largely those in receipt of the aged pension or other social security payments, are entitled to a lower co-payment. Thus the basic structure for hospital admissions and medical services has been established so that most Australians can use essential health care without financial barriers.

Private inpatient treatment is also subsidized through the reimbursement of in-hospital medical costs (at 75% of schedule fees). Although the government, in effect, sets a floor price for medical services, medical practitioners are free to set their own fees. For out of hospital medical services any fee above the bulk billing reimbursement rate cannot be covered by private health insurance and must be met by the patient as an out-of-pocket payment. For private inpatient medical services, some portion of fees in excess of the government reimbursement can be covered by private health insurance, which is limited to covering private hospital treatment (in a public or private hospital) and some ancillary services, including dentistry. Private insurance offers the choice of doctor (public patients are treated by doctors paid by the hospital); and lower waiting times particularly for elective surgery. The extent to which private patients in both public and private hospitals face out-of-pocket costs depends on their insurance policy and agreements between insurers and providers.

Bulk billing rates overall have increased over time, from under 50% of all items in 1984/1985 to around 70% in 2001; and are higher in general practice, around 75% in 2001 (DoHA statistics). Bulk billing varies by provider specialty, whether a city or rural location, and socio-economic characteristics of the area of residence. As a result, whether or not a patient faces an out-of-pocket cost for a medical service depends on provider behaviour. While there has been an expectation that concession cardholders would be bulk billed, there has been no explicit incentive to do so in the period under consideration. Although entry into the market is highly regulated, medical practitioners are free to choose the location of their practice. For general practice consultations, bulk billing rates depend on the local market for providers. Greater competition between providers increases the rate of bulk billing; higher average incomes in the local area decrease the bulk billing rate [10,11].

3. Methods and data

The degree of inequality in use of health care can be measured using the concept of a concentration curve as shown in Fig. 1. It plots the cumulative distribution of use as a function of the cumulative distribution of the population ranked by income. A distribution is equal if the cumulative distribution coincides with the diagonal. If the curve lies above (below) the diagonal this indicates that use is more concentrated among the poor (rich). A concentration index (CI) measures the degree of inequality in use as twice the area between the concentration curve and the diagonal. When it is positive, it indicates pro-rich inequality, and when it is negative, it indicates pro-poor inequality.

We would expect overall health care use to be pro-poor because lower income groups generally have poorer health status and therefore higher needs for care. A more appropriate measure of inequality would adjust

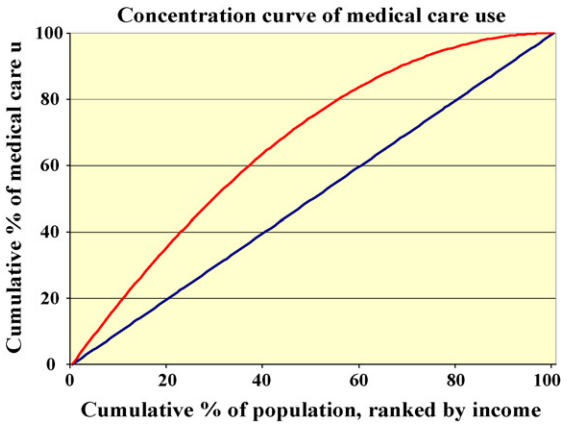

Fig. 1. Concentration curve of medical care use.

use for differences in need. In this paper we estimate the degree of inequality in need-standardized use.

The methods we use to describe and measure the degree of horizontal inequity in health care delivery are conceptually identical to the ones used in Wagstaff and Van Doorslaer (2000a) and Van Doorslaer et al. [6,12,13]. This involves the comparison of the observed distribution of medical care by income with the distribution of need. In order to statistically adjust needs for groups or individuals, the average relationship between need and use estimated for the population as a whole is used as the "norm" of what treatment is needed. We investigate the extent to which there are any systematic deviations from this norm by income level. The approach is to model health care use employing OLS regression techniques.

Health care utilisation data such as physician visits are known to have skewed distributions with typically a large majority of survey respondents reporting zero or very few visits and only a very small proportion reporting frequent use. Because these features cause violations of the standard OLS model, various specifications of non-linear, two-part models have been proposed in the literature, distinguishing between the probability of positive usage and the amount of usage conditional on use in the reference period (see [14], for a review). While these models have certain advantages over OLS specifications, their intrinsic non-linearity makes a direct application of the (linear) decomposition method described below impossible. In order to restore the mechanics of the decomposition, one has to revert to a re-linearization of the models using approximations (see [8], for an example). However, Van Doorslaer et al. [6,13] have shown that the measurement of horizontal inequality is not sensitive to the specification. We are confident that our results are not conditional on the choice of the linear standardization model.

3.1. Measurement of inequity

We first simply describe distributions of actual and need-adjusted use of health care by income quintile, each representing 20% of the total population ranked by household equivalent income from poorest to richest. "Need-expected" health care use is computed by regressing medical care use, y_i (e.g. doctor visits or hospital nights) on a set of explanatory variables

$$y_i = \alpha + \beta \ln \mathrm{inc}_i + \sum_k \gamma_k x_{k,i} + \sum_p \delta_p z_{p,i} + \varepsilon_i \quad (1)$$

We distinguish between three types of explanatory variables: the (logarithm of) the household income of individual i ($\ln \mathrm{inc}_i$), a set of k need indicator variables (x_k) including demographic and morbidity variables, and p non-need variables (z_p). α, β, γ_k and δ_p are parameters and ε_i is an error term.

Eq. (1) is used to generate need-predicted values of use, $\hat{y}_i^X$, i.e. the expected use of medical care of individual i on the basis of his/her need characteristics. It indicates the amount of medical care s/he would receive if s/he had been treated as others with the same need characteristics (e.g. age) on average. This acts as the implicit norm in the analysis: on average, individuals have to be treated similarly as others with the same need characteristics. Combining estimates of the coefficients in Eq. (1) with actual values of the x_k variables and sample mean values of the $\ln \mathrm{inc}_i$ and z_p variables, we can obtain the need-predicted, or "x-expected" values of utilisation, $\hat{y}_i^X$ as

$$\hat{y}_i^X = \hat{\alpha} + \hat{\beta} \ln \mathrm{inc}^m + \sum_k \hat{\gamma}_k x_{k,i} + \sum_p \hat{\delta}_p z_p^m \quad (2)$$

Estimates of the (indirectly) need-standardized utilisation, $\hat{y}_i^{IS}$, can be obtained as the difference between actual and x-expected utilisation, added to the sample mean (y^m)

$$\hat{y}_i^{IS} = y_i - \hat{y}_i^X + y^m \quad (3)$$

The quintile means of these indirectly standardized values provide our need-standardized distributions of medical care. They are interpreted as the expected distributions if need were equally distributed across quintiles.

Quintile distributions are difficult to compare across types of care. It is therefore useful to summarize the degree of inequality observed using a concentration index which can be computed using a simple covariance formula, as shown below for weighted data

$$C = \frac{2}{y^m}\sum_{i=1}^{n} w_i(y_i - y^m)(R_i - R^m) = \frac{2}{\mu} cov_w(y_i, R_i) \quad (4)$$

where y^m is the weighted sample mean of y, cov_w denotes the weighted covariance and R_i is the (representatively positioned) relative fractional rank of the ith individual, defined as

$$R_i = \tfrac{1}{n}\sum_{j=1}^{i-1} w_j + \tfrac{1}{2} w_i \quad (5)$$

where w_i denotes the sampling weight of the ith individual and the sum of w_i equals the sample size (n).

Testing for differences between concentration indices requires confidence intervals. Robust estimates for C and its standard error can be obtained by running the following convenient (weighted least squares) regression of (transformed) y on relative rank

$$\frac{2\sigma_R^2}{y^m} y_i = \alpha_1 + \beta_1 R_i + \varepsilon_{1,i} \quad (6)$$

where σ_R^2 is the variance of R_i and $\hat{\beta}_1$ is equal to C, and the estimated standard error of $\hat{\beta}_1$ provides the estimated standard error of C.

The concentration index of actual medical care use measures the degree of inequality while the concentration index of the need-standardized use (which is our horizontal inequity index HI) measures the degree of needs-adjusted inequality. It is worth emphasizing that coinciding concentration curves for need and actual use provide a sufficient but not a necessary condition for horizontal equity. Even with crossing curves, one could have zero inequity if, for example, inequity favoring the poor in one part of the distribution exactly offsets inequity favoring the rich in another.

3.2. Use of decomposition to explain inequality

It is possible to estimate the "contributions" of the various determinants and their relative importance. Using the regression coefficients γ_k (partial) elasticities of medical care use with respect to each determinant k can then be defined as

$$\eta_k = \frac{\gamma_k x_k^m}{y^m} \quad (7)$$

where y^m is the (population weighted mean) of y and x_k^m is the (population weighted) mean of x_k. These elasticities denote the percentage change in y result from a percentage change in x_k.

It has been shown [15] that the total concentration index can then be written as

$$C = \eta_r C_{\ln \mathrm{inc}} + \sum_k \eta_k C_{x,k} + \sum_p \eta_p C_{z,p} + GC_\varepsilon \quad (8)$$

where the first term denotes the partial contribution of income inequality, the second the (partial) contribution of the need variables, and the third the (partial) contribution of the other variables. The last term is the generalized concentration index of the error term ε. In other words, estimated inequality in predicted medical care use is a weighted sum of the inequality in each of its determinants, with the weights equal to the medical care use elasticities. The decomposition also makes clear how each determinant k's separate contribution to total income-related inequality in health care demand can be decomposed into two meaningful parts: (i) its impact on use, as measured by the use elasticity (η_k), and (ii) its degree of unequal distribution across income, as measured by the (income) concentration index (C_k). This decomposition method allows us to separate the contributions of the various determinants, and also to identify the importance of each of these two components within each factor's total contribution.

3.3. Data and variables used

We use data from the 2001 Australian National Health Survey which is conducted using a representative sample of the non-institutionalized residential population. In this study we focus on the adult sample (>18). Measurement of health care utilisation is based on a combination of various survey questions. The probability of reporting at least one visit to any

doctor (either a general practitioner or a medical specialist) in the last 12 months was derived from answers to the question "When was the last time you visited a doctor?". The actual number of visits to a general practitioner or a medical specialist was recorded only for the 2 weeks preceding the interview. We also use the probability of having been admitted to a hospital in the last 12 months, and we separately investigate admissions as a public or as a private patient.

The measurement of health as a proxy for care need was based on four types of questions. First, respondents' rating of their general health status based on five categories ('excellent', 'very good', 'good', 'fair', and 'poor'). Second, data indicating whether a respondent's health status got somewhat worse, or much worse, over the last year was used to create a dummy variable indicating deterioration in health. Third, for the analysis of hospital use, dummy variables were included for health care conditions that were constructed from detailed data on long-term conditions. Finally, nine age and gender dummies were included to reflect the fact that the meaning of the above health indicators tends to differ by age and sex.

This is a more extensive set of explanatory variables than the one used in [13]. It is well known that the inclusion of additional health information in the need standardization procedure tends to result in less pro-poor results (see [4]). This appears to be because the poor suffer from health problems more frequently and their health problems tend to be more severe. Less extensive use of health information in the need standardization process, because of the selection of a common core set of indicators for cross-country comparisons, may lead to an overestimation of pro-poor patterns. We therefore expect less pro-poor findings in this study.

Household income per equivalent adult was only available for income decile groups. Our income measure was defined as the logarithm of the mean household weekly equivalent income for the decile in which the individual's household was located (ranging from $110 for the bottom decile to $1420 for the top decile). Other explanatory variables used in the analysis included education and employment status, as these factors may affect an individual's general propensity to consume health care, but cannot often directly be influenced by health policy makers. A dummy variable for whether the respondent usually speaks English at home was included as an indicator of non-financial access to health care, and dummies for location in major city or inner region to capture differences in geographical access to doctor and hospital services.

Health insurance coverage was measured by two variables: whether the person has private health insurance coverage, and whether the person has a health care concession card.

4. Results

We present quintile distributions and concentration indices for all types of care and for both unadjusted (actual) distributions and for need-adjusted (standardized) probabilities of use. Next, we investigate the contributions of various determinants of health care use using the decomposition results.

4.1. Quintile distributions and indices

The level of health care use in Australia is quite high compared to most other OECD countries; 87% of the population has seen a doctor in the last year (28% even in the last 2 weeks) and over 14% have been admitted to a hospital within the previous year. But such comparisons of admission rates may be contaminated by uncertainties regarding the appropriate labeling of day admissions as proper admissions or outpatient visits [13].

Table 1 presents probabilities and concentration indices for all subcategories of doctor and hospital use. For all types of use except private hospital admissions, lower income groups are more intensive users, as can be seen from the negative gradients by quintile and the (significantly) negative concentration indices of actual use. This is not an unfamiliar pattern in countries with universal coverage systems and reflects the fact that the need for care tends to be inversely related to income (cf. [8]). But the degree clearly varies by type of visit and by type of admission, as is illustrated graphically in Figs. 2 and 3. Fig. 2 shows that the quintile distributions of GP visits are more concentrated among the lower income groups than the visits to medical specialists. And similarly, Fig. 3 demonstrates that the distributions of hospital admissions as public or private patient differ dramatically by income: the probability of being admitted as a public patient is three times higher than

Table 1
Fractions of users, by quintile, by type of care (NHS 2001)

	Quintile					Total	Conc/HI Index	*t*-stat
	1	2	3	4	5			
Actual use								
Any doctor, in 12 months	0.895	0.880	0.833	0.830	0.832	0.854	**−0.0164**	−6.97
A GP, in 2 weeks	0.339	0.304	0.216	0.193	0.179	0.245	**−0.1411**	−14.45
A specialist, in 2 weeks	0.079	0.068	0.057	0.055	0.065	0.065	**−0.0505**	−2.37
Any doctor, in 2 weeks	0.373	0.335	0.242	0.222	0.217	0.277	**−0.1233**	−13.7
Need standardized use								
Any doctor, in 12 months	0.852	0.850	0.847	0.853	0.864	0.854	0.0029	1.29
A GP, in 2 weeks	0.281	0.268	0.235	0.224	0.220	0.245	**−0.0535**	−5.73
A specialist, 2 weeks	0.057	0.059	0.063	0.065	0.078	0.065	**0.0634**	3.03
Any doctor, in 2 weeks	0.309	0.297	0.262	0.255	0.260	0.276	**−0.0394**	−4.6
Admission probability to: actual use								
Any Hospital, in 12 months	0.176	0.173	0.136	0.109	0.108	0.139	**−0.1205**	−8.73
Public Admit., in 12 months	0.130	0.115	0.070	0.046	0.038	0.079	**−0.2632**	−13.82
Priv Admit., in 12 months	0.046	0.058	0.066	0.063	0.070	0.061	**0.0650**	3.04
Need standardized use								
Any Hospital, 12 months	0.145	0.156	0.140	0.122	0.135	0.139	**−0.0329**	−2.47

Note: a positive/negative CI indicates a pro-rich/pro-poor distribution. Index in bold typeface indicates statistically significant difference from zero at 95% confidence level.

that of being admitted as a private patient in the bottom quintile. The reverse is true at the high end: in the top income quintile the private admission probability is twice the public admission probability.

The results for needs-standardized use are also shown in Table 1. The pattern is more mixed. The annual probability of seeing a doctor is fairly equally distributed. It is highest in the top quintile but the differences are very small and the non-significant concentration index suggests an equitable distribution. All other indices are significantly different from zero, suggesting either inequity favoring the poor (if negative) or inequity favoring the rich (if positive). GP visit probability in the last 2 weeks is pro-poor while specialist visit probability is pro-rich. The combined doctor visit probability in the last 2 weeks is also pro-poor but this is not surprising as GP visits constitute the great bulk of 2-week visits.

A similar pattern is observed when we examine hospital admission probabilities. The overall picture is one

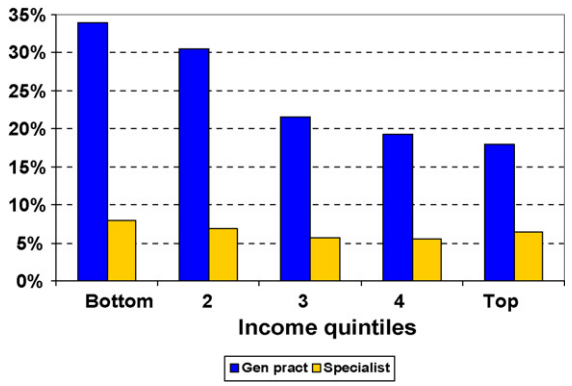

Fig. 2. GP and specialist visit probabilities in the last 2 weeks by doctor type and by quintile of equivalent income (NHS 2001).

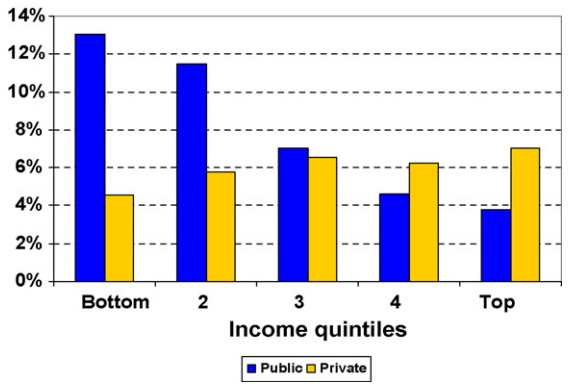

Fig. 3. Public and private admission probabilities in the last 12 months by quintile of equivalent income (NHS 2001).

Table 2
Decomposition analysis of income-related inequality in specialist visit probability

Variable	Mean	Conc Index	Regression coefficient	*t*-Value	Contrib. to Conc. Ind.	Summed contrib.
(At least one) specialist visit (actual)	0.065	−0.050				−0.050
(At least one) specialist visit (pred)	0.065	−0.114				−0.056
Horizontal Inequity Index						0.0633
Income, Log (hh income)	6.054	0.067	0.002	0.51	0.014	0.014
Health care need proxies						
Self-assessed health very good	0.325	0.110	0.005	0.98	0.003	
Self-assessed health good	0.309	−0.034	0.015	2.5	−0.002	
Self-assessed health fair	0.139	−0.217	0.050	5.17	−0.023	
Self-assessed health poor	0.050	−0.427	0.095	5.02	−0.031	−0.054
Health worse than a year ago	0.113	−0.110	0.046	4.65	−0.009	
Health much worse than a year ago	0.020	−0.236	0.138	4.51	−0.010	−0.019
Male 34–49	0.160	0.181	−0.005	−0.76	−0.002	
Male 50–64	0.115	0.070	−0.005	−0.53	−0.001	
Male 65–74	0.049	−0.384	0.015	0.94	−0.004	
Male 75+	0.032	−0.414	0.009	0.46	−0.002	−0.009
Female15–34	0.146	0.074	0.045	3.79	0.008	
Female 34–49	0.159	0.066	0.011	1.33	0.002	
Female 50–64	0.108	−0.047	0.016	1.4	−0.001	
Female 65–74	0.050	−0.425	0.009	0.54	−0.003	
Female 75+	0.041	−0.487	−0.028	−1.65	0.009	
Fem 20–35 with child(ren)	0.090	−0.076	−0.011	−0.8	0.001	0.015
Other infectious diseases	0.010	−0.054	0.057	1.87	−0.001	
Other Neoplasms	0.003	−0.009	−0.014	−0.35	0.000	
Diabetes with no complications	0.035	−0.276	0.040	2.2	−0.006	
Type I diabetes mellitus	0.006	−0.215	0.086	1.85	−0.002	
Other endocrine disorders	0.087	−0.158	−0.004	−0.37	0.001	
Peptic ulcer, hemorrhage, other specified gastrointestinal disorders	0.036	−0.211	−0.005	−0.32	0.001	
Other gastrointestinal disorders	0.074	−0.097	−0.005	−0.5	0.001	
Rheumatoid arthritis	0.032	−0.299	0.024	1.38	−0.004	
Osteoporosis	0.030	−0.236	0.046	2.37	−0.005	
Other musculoskeletal disorders	0.404	−0.082	0.004	0.71	−0.002	
Disorders of immunity	0.003	0.030	0.106	1.41	0.000	
Iron deficiency and anemias	0.016	−0.069	0.005	0.27	0.000	
Drug/alcohol abuse	0.008	−0.122	−0.004	−0.14	0.000	
Personality disorders	0.048	−0.169	0.028	1.88	−0.003	
Depression	0.007	−0.227	−0.035	−1.16	0.001	
Anxiety disorders	0.051	−0.121	0.001	0.08	0.000	
Other psychiatric disorders	0.034	−0.261	0.029	1.65	−0.004	
Other developmental disability	0.001	−0.049	−0.040	−0.97	0.000	
Seizure disorders and convulsions	0.007	−0.216	0.060	1.43	−0.001	
Mononeuropathy/injuries	0.076	0.007	0.005	0.53	0.000	
Hypertensive heart disease	0.142	−0.227	0.006	0.66	−0.003	
Other heart disease	0.001	−0.218	−0.097	−3.44	0.000	
Cerebrovascular disease	0.028	−0.407	0.009	0.5	−0.002	
Vascular disease	0.015	−0.373	−0.011	−0.51	0.001	
Other circulatory disease	0.040	−0.162	−0.003	−0.19	0.000	
Chronic obstructive pulmonary disease	0.010	−0.417	−0.047	−1.84	0.003	
Asthma	0.106	−0.026	0.001	0.11	0.000	
Other lung disorders	0.032	−0.185	−0.002	−0.11	0.000	
Glaucoma	0.013	−0.257	0.048	1.65	−0.002	

Table 2 (*Continued*)

Variable	Mean	Conc Index	Regression coefficient	*t*-Value	Contrib. to Conc. Ind.	Summed contrib.
Cataract	0.029	−0.362	0.038	1.86	−0.006	
Other eye disorders	0.654	−0.031	0.010	1.69	−0.003	
Hearing loss	0.141	−0.141	0.013	1.73	−0.004	
Other ear, nose, throat disorders	0.281	0.033	−0.003	−0.59	0.000	
Urinary obstruction and retention	0.026	−0.144	0.003	0.15	0.000	
Incontinence	0.012	−0.320	0.025	0.88	−0.001	
Other urinary tract disorders	0.019	−0.019	0.017	0.87	0.000	
Other dermatological disorders	0.034	0.006	−0.015	−1.27	0.000	
Major symptoms, abnormalities	0.004	−0.237	0.122	1.87	−0.002	
Minor symptoms, signs, findings	0.103	−0.072	0.020	2.22	−0.002	−0.047
Other socioeconomic and demographic characteristics						
Education: year 9–11 completed	0.385	−0.012	0.004	0.58	0.000	
Education: year 12 or more completed	0.380	0.220	0.009	1.25	0.012	0.012
Self-employed	0.100	0.066	0.007	0.99	0.001	
Unemployed	0.034	−0.613	−0.019	−1.9	0.006	
Out of labor force	0.344	−0.401	0.019	2.35	−0.041	−0.034
Main language spoken not English	0.150	−0.139	−0.017	−2.46	0.005	0.005
Resident of major city	0.667	0.048	0.025	4.18	0.012	
Resident of inner region	0.213	−0.098	0.017	2.47	−0.006	0.007
Health insurance status variables						
Private insurance	0.516	0.207	0.024	4.54	0.039	0.039
Concession card	0.356	−0.507	−0.005	−0.71	0.015	0.015

of a very pro-poor distribution, which remains slightly pro-poor even after standardization for need, but this obscures two very different patterns underlying admission as either a public or as a private patient. It is, of course, not possible to estimate the need for a public or a private admission separately from these data, only the general need for a hospital admission, which is why no need standardized distributions are presented separately for public and private patient admissions.

4.2. Decomposition of specialist visits and public/private admissions

A factor contributes to inequality in use if it is both distributed unequally by income and has an effect on the probability of using health care. Contribution is not meant in a causal sense here but means that it helps to explain the association between use and income rank through its (partial) association with use of care. The mechanics of the decomposition method are illustrated in Table 2 for the probability of a specialist visit. The table presents the mean, concentration index and regression coefficient for each explanatory variable in the model. These are translated into an inequality contribution using Eq. (7) in Section 3.2. The CI for actual specialist visits (−0.050) corresponds to that presented in Table 1. The HI is derived as the CI minus the sum of the contributions of all the need indicators. It can be seen that virtually all the contributions of need proxies (especially self-assessed health and chronic conditions) are negative because morbidity is more concentrated among the lower income groups and increases the likelihood of using specialist services.[1]

More interesting are the contributions of the non-need determinants, i.e. the factors affecting the difference between the actual distribution and the one expected on the basis of need indicators alone. It can be seen that the contribution of the (partial) effect of income itself is much smaller than that of, for instance, not having paid employment. Because private insurance is associated both with higher income and higher

[1] Unlike Jones et al. [11], we have not bootstrapped the entire decomposition procedure in order to estimate standard errors for the contributions. Hence, we can only perform statistical inference on the regression coefficients, not on the full decomposition terms in Eq. (7).

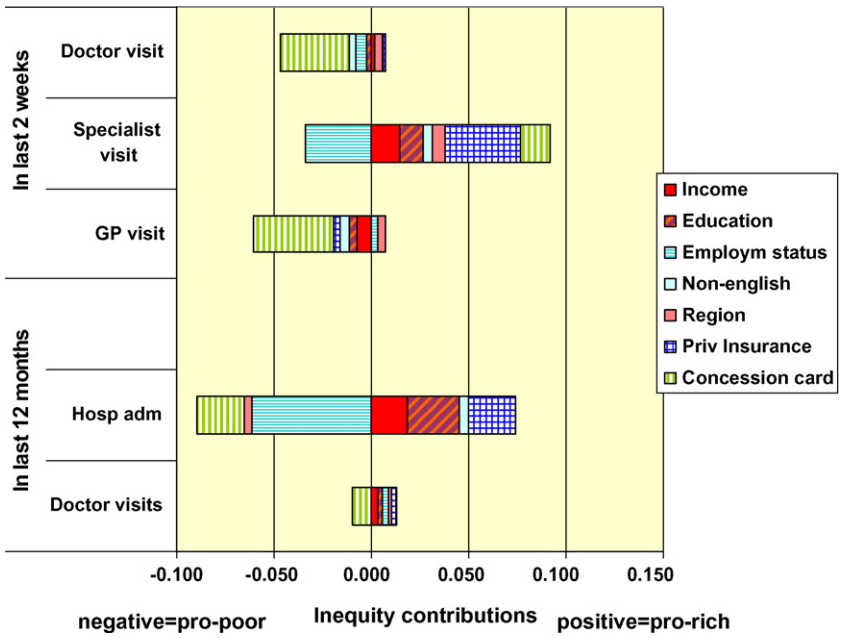

Fig. 4. Decomposition of inequity into contributions of covariates. *Note*: The bars represent the (summed) inequity contributions of the non-need variables taken from the last column of Table 2.

use of specialist visits, it has a strong positive contribution to the observed pro-rich distribution. Interestingly, for the opposite reasons, i.e. lower income associated with lower use, ceteris paribus, the concession card also contributes positively to specialist use inequality. These findings should be interpreted with caution. Obviously, private insurance and concession card status may be endogenous and it is likely that selection into these states is at least partially responsible for their contributions.

The results of the decomposition exercise for all types of care are presented graphically in Fig. 4, which shows the contributions to the measured degree of inequity after subtracting the need contributions following Eq. (8). The most important contributing factors are the insurance and employment status variables. Concession cards contribute substantially to the observed pro-poor distribution of the GP and all doctor visit probability. This is most likely due to the propensity for GPs to bulk bill concession cardholders, resulting in zero out-of-pocket costs to patients. In all cases, private insurance shows a positive contribution to income-related inequity, but only for the probability of a specialist visit or a hospital admission, this contribution is substantial. Less obvious is the role played by employment status, which reduces inequity in specialist use and makes hospital admission pro-poor. To some extent, this may be because, all else equal, not being in paid employment may be an indicator of need for care. If, for instance, most of the contribution is related to disability status, then in fact the contribution may be largely need-driven. If this were true, and employment status could be seen as a proxy for care need, then inclusion of its contribution among the need factors and its deduction from overall inequity, would make specialist even more pro-rich and would turn hospital admissions from pro-poor into pro-rich inequity.

5. Conclusion

Overall, based on our analysis of the NHS 2001 data, we can conclude that Medicare largely does appear to be attaining an equitable distribution of health care access: Australians in need for care do get to see a doctor and to be admitted to a hospital. However, they are not equally likely to see the same doctor and to end up in the same hospital bed. As in other OECD countries,

while lower income Australians are more likely to consult a general practitioner, those on higher incomes are more likely to consult a specialist, all else equal. Those on higher incomes are also more likely to be admitted into hospital as a private than as a public patient. The unequal distribution of private insurance by income further stimulates the phenomenon that the better-off and the less well-off do not receive the same mix of services. While the partial effect of private insurance identified in this paper cannot be interpreted as causal because it is at least partially endogenous, European panel data evidence suggests that controlling for endogeneity does not necessarily reduce the specialist utilisation effect of (supplementary) private insurance. Indeed, using employer-provided cover as an instrument, Jones et al. [16] find that there is positive (rather than adverse) selection into such coverage and that correction for this selection effect even increases the utilisation effect. So the estimated contribution of private insurance need not necessarily be an upper bound. For Australia, Doiron et al. [17] also find a positive selection effect of SAH on the demand for private insurance cover, but show that this effect may be due to individual traits related to risk attitudes reflected in SAH reporting. They recover the standard (negative) adverse selection when using more objective indicators like the number of long-term conditions.

Similarly, the chances of being admitted as a public or private patient are very unequally distributed and very much related to the unequal distribution of private insurance and medical cards by income. Whether or not the observed unequal care mix by income is undesirable probably depends on the extent to which it translates into unequal health outcomes. Evidence from other countries suggests that unequal access to specialist care services may result in unequal outcomes for cardiac patients (e.g. Alter et al. 2004, [18] for Ontario), but such evidence is currently lacking for Australia.

Our study was somewhat constrained by the limitations of the NHS 2001 survey data. As is true for all surveys, all information – on both income, health and use – is self-reported and therefore may to some degree be subject to inaccurate recall, or misreporting. In addition, and because of the typical infrequency of health care utilisation, the very short recall period of 2 weeks makes it less than ideal for a more detailed analysis of its appropriateness. And finally, it is unclear to what extent the hospital admission data used here excludes some same-day admissions, which is not unimportant because they constitute around half of all hospital admissions in Australia. Most of these limitations could be overcome if administrative health care data such Medicare records could be linked with measures of socioeconomic status such as income from taxation records. This is something we believe deserves serious consideration in future research.

These findings have important policy implications. First, if the more rapid future growth in private hospital capacity continues, lower income groups are likely to face less equitable access to hospital care. Low-income groups are less likely to use specialist medical services, where co-payments are higher. Further rises in co-payments for specialists may widen this gap, and increasing co-payments for general practitioners may reduce the pro-poor inequity observed in 2001. Since 2001 the Government has introduced financial incentives to bulk-bill GP consultations for concession card holders and children. However, these incentives have not been directed to specialists, where there are significant pro-rich inequalities. A new Medicare Safety Net introduced in 2004, designed to limit out-of-pocket payments for medical services, could be expected on the basis of these findings to target specialists services and favor high-income groups. Indeed, Van Gool et al. [19] have concluded that this has occurred. Further from 2007 private health insurers will be allowed to provide insurance cover for out of hospital care, including chronic disease management and prevention programs [20]. Our findings suggest that private health insurance already contributes to the differential use of health services across income groups. These new developments may reduce the overall pro-poor equity of the Australian health care system. While future research is needed, this study provides a valuable base to allow monitoring of subsequent changes in access.

The study also has some international implications. In particular it has shown overall equity can be achieved in countries where there is a large private sector, through a universal public system, which compensates for inequities in access to private care. Given the increasing trend towards private delivery in many countries, it is important to understand the equity implications of policies that affect the public private mix.

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